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# A note on the competing causes of high inflation in Bulgaria during the 1990s: money supply or exchange rate?\*

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## Abstract

This note aims at analyzing Bulgaria's high inflation regime during the 1990s. Two competing causes of high inflation are explored: changes in the rate of growth of the money supply in the economy and changes in the foreign exchange rate. Both correspond to traditional theoretical explanations: the monetarist view and the balance of payments approach. Evidence suggests that a variation in the exchange rate is significant in explaining the high inflation regime in Bulgaria whereas monetary growth appears to be insignificant. Consequently, the paper underlines the importance of stabilizing the exchange rate in the short run in order to avoid high inflation.

**Key words:** Inflation, Money supply, Exchange rate

**JEL Codes:** E31, P22

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## 1. Introduction

According to monetarist economists like Friedman (1968, p.12) “every major inflation has been produced by monetary expansion”. This being so, the best way to control inflation is to control the money supply. Though hyperinflation can be viewed as a more radical phenomenon, the same logic applies in the influential work of Cagan (1956) who shows that the rate of increase in monetary growth is essential in explaining price dynamics. A development of this line of research is also found in the work of the rational expectations school which claims that budget deficits are the cause of inflation and, above all, of hyperinflation. In such circumstances, governments finance their expenditure by issuing money: put bluntly, they sell bonds to the central bank and use the newly printed money to pay for public spending. Accordingly, monetarist approaches, notwithstanding their successive refinements and irrespective of hyperinflationary episodes, share a common characteristic: they claim an increase in the money supply precedes the rise in the level of prices.

In contradistinction to the monetarist approach, the balance of payments school focuses on war reparations, external debt and, above all, on the depreciation of the exchange rate as the main source of high inflation regimes and episodes of hyperinflation. According to Câmara and Vernengo (2001), Helfferich (1927) is “the most notorious defender” of this school of thought. In the same vein, an author like Robinson (1938) emphasizes the concept of a strong foreign exchange/wage spiral in the case of German hyperinflation. Here, plans and expectations are no longer undertaken under normal circumstances in the sense that agents start making their domestic transactions in a foreign currency, completely rejecting the domestic one. Extending the argument, Kaldor’s (1982, p. 61) description of the German situation in 1923 is particularly striking: “Everything from newspapers to railway tickets and to daily wages was ‘indexed’ to the daily market price of the US dollar (...). If the dollar remained unchanged for a day, prices and wages (...) remained stable for the day.”

Compared to the other transition economies during the 1990s, it has to be recognized that economists have paid little attention to the various origins of Bulgarian hyperinflation.<sup>1</sup> Most of the time, they underestimated or ignored the role of exchange rate variations as a valid cause in the emergence of a high inflation regime and, ultimately, of hyperinflation. Moreover, empirical studies mainly focused on the end of hyperinflation due to the implementation of a currency board in 1997 or on the fact that the maintenance of currency board arrangements may have been advantageous regarding the credibility of monetary authorities.<sup>2</sup>

The purpose of this article is to determine whether the exchange rate, as an explanatory variable for inflation, is significant in Bulgaria before the hyperinflationary episode, confirming or invalidating the arguments of the balance of payments school as sustained by Helfferich, Robinson, and Kaldor. Moreover, we try to narrow the gap between the many empirical studies in the monetarist tradition and the few in the balance of payments approach based on the exchange rate. We do not overlook the usual Post-Keynesian explanations of inflation as expounded, for example, in Davidson and Weintraub (1973) and Lavoie (2014, chapter 8). Then, according to Post-Keynesians, inflation is a consequence of the distributive conflict between wage earners and profit makers. In other words, they consider wage increases above labor productivity, changes in mark-up,

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<sup>1</sup> See Charles and Marie (2016) for an alternative definition of hyperinflation.

<sup>2</sup> See, for instance, Ivanova (2009) and Begović *et al.* (2016) on the various effects of the currency board.

demand-pull effects, and also exchange rate variations to be key variables. The reader interested in this approach and in the theoretical filiation and consistency between the balance of payments approach and Post-Keynesian analyses of inflation will find a stimulating literature review in Bastian and Setterfield (2020). Hereafter, in line with the balance of payments approach, we shall focus specifically on the role played by exchange rate variations.

In order to clarify our motivation, we test two opposing theories: the balance of payments approach in which causality runs from exchange rate to prices and the quantity theory of money in which causality runs from money supply to prices. The first approach is fully consistent with the view developed by Davidson and Weintraub (1973): by analogy, an exchange rate variation may have a similar effect to an increase in money wages, entailing inflation and growth of the money supply.

The structure of the paper is as follows. Section 2 presents a short review of the empirical literature in both traditions. Section 3 briefly deals with Bulgaria's chronology, the data utilized, and the econometric methodology while section 4 presents the empirical results. Section 5 draws some conclusions and sets out some economic policy considerations.

## 2. Literature review

The monetarist tradition has given rise in the last two decades to a large number of studies. In the case of ex-Yugoslavia, Petrović *et al.* (1999) show that excessive monetary growth was at the origin of hyperinflation. Moreover, their Granger causality tests indicate that causality runs from money supply growth to currency depreciation and from currency depreciation to inflation. For Greece, Karfakis (2002) indicates that a shock to the money supply affected prices in a similar way. Gillman and Nakov (2004) find strong causality between monetary growth and inflation for Hungary and Poland over the period 1986–2002. De Grauwe and Polan (2005), using a sample of about 160 countries, provide evidence of a link between the money supply and the price level. However, this relationship disappears when they retain a subsample of countries with monetary growth of less than 15% per annum, contrasting with the results from Frain (2004) who confirms, independently of the sample, the strong correlation between money supply growth and inflation in a multi-country analysis. In the case of China, Lee and Chien (2008) find that the long-run income elasticity for real money balances (M1 and M2) is close to one, as suggested by the quantity theory of money (QTM). Tawadros (2008) reports strong evidence in favor of QTM, even in the medium-run, for four European countries (Germany, Poland, Austria, and Hungary) during the 1920s. His results unambiguously indicate that inflation varies positively with money supply growth. Finally, Ojede (2015), using a sample of 54 developing countries, reports strong statistical evidence that inflation is mainly driven by monetary growth.

By contrast, empirical studies investigating the impact of the foreign exchange rate on prices, within the balance of payments (BP hereafter) school, are really few in number in spite of some interesting results.<sup>3</sup> For example, Pastor (1991) shows that during the high

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<sup>3</sup> As emphasized by Câmara and Vernengo (2001), there is a close relationship between neo-structuralists and the BP school: both reject the QTM and recognize that the exchange rate may have a substantial effect on price dynamics; but they also insist on the impact of distributive conflict and indexation mechanisms as underlined by Missio *et al.* (2015, p. 260). For a recent empirical assessment for Argentina see, for example, Frenkel and Friedheim (2017).

inflation regime in Bolivia (from March 1982 to August 1985) changes in the parallel exchange rate fuelled inflation. Moreover, reducing the sample to just the hyperinflationary period makes the exchange rate virtually the only explanatory variable for price dynamics. Similarly, Burdekin and Burkett (1996), examining the hyperinflationary period in Germany, elaborate a series of econometric models in which exchange rate depreciation has systematic, strong, and significant effects on inflation. Maswana (2005) finds a unidirectional Granger causality running from exchange rate to inflation for the Democratic Republic of the Congo from January 1990 to September 1996. Nevertheless, the author also finds unidirectional Granger causality running from money supply growth to the exchange rate which, at best, provides mixed results since the link between money supply and prices is indirectly maintained. More recently, Jalil *et al.* (2014), though testing for a fiscal deficit explanation of inflation in Pakistan, also find a significant effect of the exchange rate.

There are macroeconomic studies on the Bulgarian question touching on inflation among other factors (see, for example, Minassian, 1998). Nevertheless, there is a marked need for fresh evidence about this high inflation regime in the 1990s, given the absence of econometric studies on Bulgaria dealing with the BP tradition.

### **3. Chronology, data, and methodology**

Bulgaria entered the post-socialist transitional phase late in 1989. In March 1990 the government adopted a moratorium on the payment of its foreign debt and economic liberalization reforms began in 1991: prices were deregulated, planning was abandoned, free entry onto the market was sought after, as was the liberalization of foreign trade. Dobrinsky (2000, p. 583) indicates that the country adopted a strategy of exchange-rate floating with stabilization of the monetary base. Being excluded from the international financial markets, the Bulgarian authorities maintained foreign exchange controls.

Economic results were pretty poor (see Table 1); GDP plummeted over the period 1991–1993. The year 1993 was marked by a downturn in exports in a difficult international context due to the trade blockade on former Yugoslavia. It should be noted that slightly positive growth rates were observed in 1994 and 1995. These were made possible by a surge in exports in 1994 and by increased investment in 1995. Inflation, estimated from changes in the consumer price index (CPI), was very high as from 1991. While the variation in the foreign exchange rate was above inflation in 1991, it fell below inflation from 1992 to 1995 inclusive. That caused a fall in the competitiveness of Bulgarian prices or a rise in the value of the domestic currency in real terms. In 1995 and early 1996, Bulgaria resumed servicing its foreign debt. This improvement of the economic context was thwarted by the fragility of the banking and financial system, which became apparent from the end of 1995. In agreement with contemporary analyses by the IMF, we emphasize that this financial fragility is a fundamental factor in understanding Bulgaria's economic trajectory. These factors made it possible for a dangerous mechanism to be put in place: collapse of the banking system was averted or delayed by refinancing conditions facilitated by the Bulgarian National Bank (BNB), behavior that resulted in diminished confidence in the domestic currency with respect to foreign currencies at the very time when the central bank was seeking to maintain stable exchange rates. Hyperinflation, defined as the flight from domestic currency into foreign currencies, was nonetheless warded off until 1997.

**Table 1: Macroeconomic performances of Bulgaria**

|                           | 1992 | 1993 | 1994 | 1995 | 1996  | 1997   |
|---------------------------|------|------|------|------|-------|--------|
| GDP (real growth rate, %) | -7.2 | -1.4 | 1.8  | 2.8  | -9.4  | -5.5   |
| Inflation (%)             | 82.0 | 72.8 | 96.0 | 62.1 | 123.0 | 1061.2 |
| Unemployment rate (%)     | 15.3 | 15.8 | 12.4 | 11.1 | 12.5  | 13.7   |

Source: IMF and BNB, authors' calculations

The monthly time series mobilized in the article come from the BNB annual reports for the foreign exchange rate and monetary mass. The consumer price index is from the IMF's IFS (International Financial Statistics) database over the period 1991m04–1996m11. The early months of 1991 are omitted since they characterize the transition shock with the deregulation of many prices and the opening-up of the economy. Therefore we start the empirical analysis in April 1991. Such a decision is commonly adopted when dealing with transitional economies (Slavova, 2003). Regarding the end of the data sample, we choose the month of November since the hyperinflationary logic starts, strictly speaking, in December with an ever increasing inflation rate until the peak of February 1997.

Following the BP view, in high inflation regimes, the main determinant of inflation is assumed to be the variation in the exchange rate. Accordingly, the following regression is estimated for Bulgaria:

$$inf_t = \alpha_0 + \alpha_1 \hat{e}_t + \alpha_2 \hat{m}_t + \varepsilon_t \quad (1)$$

where  $inf_t$  is the rate of inflation defined as the first difference of the consumer price index logarithm,  $\hat{e}$  is the first difference of the indirect foreign exchange rate logarithm (1 U.S. Dollar for 'e' Leva),  $\hat{m}$  is the first difference of the logarithm of the monetary mass (M2), and  $\varepsilon_t$  is an error term.<sup>4</sup> Equation (1) also incorporates monetary growth as an explanatory variable in order to assess both competing theories. The parameters,  $\alpha_{i=1,2}$ , indicate the elasticity estimates of exchange rate and monetary variations, respectively. The value of  $\alpha_1$  ( $\alpha_2$ ) is expected to be high and very significant if the BP school (the QTM) prevails while an insignificant coefficient could disqualify such a view.

Nevertheless, the previous equation cannot be estimated using an ordinary least squares procedure or a VAR methodology if the data are non-stationary. So, if the time series are  $I(1)$  cointegration methods must be applied. Another difficulty appears when the data series are  $I(0)$  and  $I(1)$ . In this case, econometric theory strongly suggests the use of an Autoregressive Distributed Lag (ARDL) cointegration method to test the existence of a long-run relationship between two or more variables (Pesaran and Shin, 1999 and Pesaran *et al.*, 2001). The ARDL bounds testing approach to cointegration has several advantages over other methods. As indicated above, there is no need for all the variables to have the same order of integration. It is efficient, even in the case of a small sample, provided we use Narayan's (2005) critical bounds. Moreover, it also works with an endogenous regressor.

The ARDL bounds testing approach must be implemented in three steps before estimating equation (1). First, we need to determine the order of integration of the variables, bearing in mind that a series must not be  $I(2)$ , otherwise critical values used to

<sup>4</sup> Monetary growth is deseasonalized using the US Census Bureau's X13 seasonal adjustment program.

determine the existence of cointegration are no longer valid. Second, to implement the bounds testing procedure we estimate the following conditional ARDL model:

$$\begin{aligned} \Delta inf_t = & \alpha_0 + \sum_{i=1}^n \beta_i \Delta inf_{t-i} + \sum_{i=0}^n \gamma_i \Delta \hat{e}_{t-i} + \sum_{i=0}^n \varphi_i \Delta \hat{m}_{t-i} \\ & + \theta_1 inf_{t-1} + \theta_2 \hat{e}_{t-1} + \theta_3 \hat{m}_{t-1} + \varepsilon_t \end{aligned} \quad (2)$$

where the second part of the equation, containing the  $\theta$ s, corresponds to the long-run relationship. The bounds testing approach is based on the F-statistic and tests the null of no cointegration, i.e.  $H_0: \theta_i = 0$  with  $i = 1, 2, 3$ , against the alternative,  $H_1: \theta_i \neq 0$ . Pesaran *et al.* (2001) report two sets of critical values that provide critical value bounds for all classifications of the regressors as purely  $I(1)$ , purely  $I(0)$ , or mutually cointegrated. If the calculated F-statistic is above the upper level of the bound, the null hypothesis is rejected which indicates the presence of cointegration between the variables. Nevertheless, critical values are determined for long sample sizes of at least 500 observations. So, considering we have 68 observations, it is necessary to use this procedure with the corrected critical values provided by Narayan (2005) for reduced samples. Finally, once the bounds test is conclusive, the long-run model can be implemented from equation (2) where the first-differenced variables are all equal to zero.

#### 4. Empirical results

First of all, we need to test for the existence of unit roots to ensure that all the variables satisfy the underlying assumptions of the ARDL bounds testing approach of cointegration. The results are summarized in Table 2, using standard Augmented Dickey-Fuller (ADF) and Elliott-Rothenberg-Stock (1996) tests (ERS) with a constant. It turns out that the rate of inflation is stationary at level for each test. The exchange rate depreciation is stationary only at first difference with ERS, rejecting the possibility of being  $I(2)$ . In the same vein, an ambiguity subsists regarding monetary growth. Therefore, with a mixture of  $I(0)$  and  $I(1)$  series and/or a doubt about their stationarity, there is a strong justification for utilizing an ARDL approach in order to determine the existence of long-run relationship among the variables.

**Table 2: Unit root tests in level and first difference**

|             | ADF          |              | ERS          |              |        |     |
|-------------|--------------|--------------|--------------|--------------|--------|-----|
|             | <i>level</i> | <i>diff.</i> | <i>level</i> | <i>diff.</i> |        |     |
| $inf_t$     | -3.345**     | [0]          | 2.038**      | [0]          |        |     |
| $\hat{e}_t$ | -2.995**     | [1]          | 3.666        | [1]          | 0.247* | [2] |
| $\hat{m}_t$ | -3.179**     | [0]          | 4.090        | [0]          | 1.719* | [0] |

Notes: Critical values for the ADF test are -3.531 (1%) and -2.905 (5%) for inflation and monetary growth and -3.533 (1%) and -2.906 for exchange rate variations. Critical values for the ERS test are 1.898 (1%) and 3.020 (5%). The number of optimal lags, in square brackets, is determined according to the Schwarz information criterion with a maximum lag length of 6. Here, (\*) and (\*\*) denote significance at the 1% and 5% levels.

The next step consists in implementing the bounds test with the appropriate critical values for small sample sizes. We retain the Schwarz information criterion with a maximum lag length of 6 for the dependent variable and the regressors and find an ARDL (1,1,0) model. The second bounds test concerns exclusively the relation between inflation and exchange rate variations, the econometric procedure selects an ARDL (1,1). In both cases, the F-statistic for the bounds test is largely above the upper bound even at the 1% level. Consequently, we reject the null hypothesis of no long-run relationship between the variables.

**Table 3: Bounds test for the existence of cointegration**

| F-statistics                                   | CV 5%        |              | CV 1%        |              |
|--|--------------|--------------|--------------|--------------|
|  | <i>I</i> (0) | <i>I</i> (1) | <i>I</i> (0) | <i>I</i> (1) |
| ARDL (1,1,0) $inf_t = f(\hat{e}_t, \hat{m}_t)$ |              |              |              |              |
| 15.476   | 3.243        | 4.043        | 4.398        | 5.463        |
| ARDL (1,1) $inf_t = f(\hat{e}_t)$              |              |              |              |              |
| 20.891   | 3.780        | 4.327        | 5.157        | 5.957        |

Notes: critical values come from Narayan (2005, Appendix, case II)

Next, we estimate the long and short-run determinants of inflation. Our results are shown in Table 4 in the column headed Model I. Note that the coefficient of exchange rate variation is high and strongly significant as an explanatory variable for inflation. In contrast money supply growth is insignificant over the period. Accordingly the BP school seems to prevail over the QTM tradition in the case of Bulgaria. A 1% increase in exchange rate depreciation induces a 0.71% increase in inflation. The error correction term (ECT) is negative and statistically significant at the 1% confidence level, indicating that a deviation from the long-run equilibrium between variables is corrected for in each period to return to this equilibrium. The value  $-0.579$  shows that it takes a little less than two months to return to the long-run equilibrium. As for the short-run, results remain similar since the exchange rate (actual and lagged) is significant while money growth is insignificant in explaining inflation. In Model II, given the absence of the effect of money growth, we implement exactly the same procedure with the exchange rate as the sole explanatory variable. As shown in the second column of Table 4, the results are roughly the same: a 1% increase in exchange rate depreciation still causes a 0.71% rise in inflation. This emphasizes the crucial role of exchange rate variations in triggering the high inflation regime that prevailed in Bulgaria during the 1990s before the hyperinflationary episode.

In order to assess the strength of these error correction models, we also perform several diagnostic tests in the third part of Table 4 which indicates p-values of chi-squares. We find no evidence of serial correlation in the residuals based on the Lagrange multiplier test, an absence of heteroskedasticity through the White test, and a correct functional form. Regarding the normality test, we need to be cautious because of our reduced sample. Indeed, when the Jarque-Bera test is performed on small samples there may be considerable size distortion leading to the normality of residuals being rejected even when valid (see Thadewald and Buning, 2007). In order to overcome this difficulty, we provide a bootstrapped p-value for the JB test in which the residuals from the ARDL model

were bootstrapped 5000 times. In both cases, we observe unambiguously that the residuals are normally distributed.

**Table 4: Estimated coefficients from the ARDL model**

|  | Model I             | Model II            |
|--|---------------------|---------------------|
| <b>Long-run determinants of inflation</b>  |                     |                     |
| $\hat{e}_t$                                | 0.713*<br>(7.049)   | 0.712*<br>(7.222)   |
| $\hat{m}_t$                                | -0.081<br>(-0.252)  | NA<br>NA            |
| <i>Constant</i>                            | 0.030**<br>(2.190)  | 0.027*<br>(4.205)   |
| $ECT_{t-1}$                                | -0.579*<br>(-7.869) | -0.584*<br>(-8.041) |
| <b>Short-run determinants of inflation</b> |                     |                     |
| $inf_{t-1}$                                | 0.425*<br>(4.647)   | 0.415*<br>(5.021)   |
| $\hat{e}_t$                                | 0.195*<br>(4.394)   | 0.196*<br>(4.477)   |
| $\hat{e}_{t-1}$                            | 0.214*<br>(3.739)   | 0.219*<br>(4.106)   |
| $\hat{m}_t$                                | -0.046<br>(-0.257)  | NA<br>NA            |
| <i>Constant</i>                            | 0.017**<br>(2.281)  | 0.016*<br>(3.279)   |
| $R^2$                                      | 0.710               | 0.710               |
| <b>Diagnostic test</b>                     |                     |                     |
| Serial correlation                         | 0.616               | 0.630               |
| Heteroskedasticity                         | 0.373               | 0.535               |
| Normality                                  | 0.456               | 0.459               |
| Functional form                            | 0.166               | 0.149               |

Notes: (\*) and (\*\*) denote 1 and 5% level of significance, t-statistics for coefficients are in ( ). For the diagnostic test, the results stand for the p-values.

At this point, we can check for the presence of cointegration between exchange rate depreciation and money growth as an alternative or indirect version of the QTM. Following this logic, too rapid a rise in the domestic money supply would lead to sharper exchange rate depreciation and, lastly, to an increase in prices leaving inflation still

dependent on the quantity of money in the economy. In Table 5, the calculated F-statistic based on an ARDL (2,0) is below the lower bound at the 5% level (but also at the 10% level). Here, we cannot reject the null hypothesis of no long-run relationship between both variables. Moreover, the model passes all the diagnostic tests, we find no evidence of serial correlation in the residuals, an absence of heteroskedasticity, normally distributed residuals and a correct functional form at the standard 5% level of significance. Consequently, this result also seems to contradict the alternative version of the QTM in which the positive link between monetary growth and inflation is maintained indirectly through variations in the exchange rate.

**Table 5: Bounds test for the existence of cointegration between  $\hat{e}_t$  and  $\hat{m}_t$**

| F-statistics   | CV 5%    |       | CV 1% |       |
|--|----------|-------|-------|-------|
|  | I(0)     | I(1)  | I(0)  | I(1)  |
| ARDL (2,0) $\hat{e}_t = f(\hat{m}_t)$<br>3.047           | 3.780    | 4.327 | 5.157 | 5.957 |
| Short-run determinants of the variation in exchange rate |          |       |       |       |
| $\hat{e}_{t-1}$  | 0.218    |       |       |       |
|  | (1.677)  |       |       |       |
| $\hat{e}_{t-2}$  | 0.351**  |       |       |       |
|  | (2.734)  |       |       |       |
| $\hat{m}_t$  | -0.292   |       |       |       |
|  | (-0.647) |       |       |       |
| Constant   | 0.030    |       |       |       |
|  | (1.479)  |       |       |       |
| $R^2$  | 0.216    |       |       |       |
| Diagnostic test  |          |       |       |       |
| Serial correlation                                       | 0.223    |       |       |       |
| Heteroskedasticity                                       | 0.682    |       |       |       |
| Normality  | 0.190    |       |       |       |
| Functional form  | 0.062    |       |       |       |

Notes: Critical values come from Narayan (2005, Appendix, case II). (\*) and (\*\*) denote 1 and 5% level of significance, t-statistics for coefficients are in ( ). For the diagnostic test, the results stand for the p-values.

These empirical results confirm that the BP approach should be considered more thoroughly when dealing with high inflation levels. We show that changes in the exchange rate were already significant in Bulgaria even before the dramatic hyperinflationary episode of February 1997. Moreover, the explosion in prices was not the consequence of money growth but rather a rejection of the domestic currency in favor of certain foreign currencies. This violent phenomenon of currency substitution can be explained by at least two factors.

First, the uncontrolled deregulation of the banking sector increased financial fragility. For instance, Berlemann and Nenovsky (2004) explain that the appearance of Ponzi type financial pyramids was facilitated by the very liberal policy of the Bulgarian National Bank in granting banking licenses. Poirot (2003, pp. 41-47) claims the financial fragility was the consequence of generalized soft budget constraints compounded by a culture of “non-payment” in the banking sector facilitated by inadequate compliance with contracts and inefficient supervisory measures. Consequently, agents progressively lost their trust in the banking system and tried to protect their savings by replacing Leva with US Dollars and Deutschmarks. Second, this rejection of the domestic currency was largely magnified by the appearance of a price/exchange rate indexation mechanism as underlined long ago by Robinson (1938) and Kaldor (1982) in the case of German hyperinflation.

## 5. Conclusion

In the present study we compare two competing theories to explain the high inflation regime in Bulgaria during the 1990s: the QTM tradition and the BP school. We test the relationships among inflation, exchange rate variations, and money supply growth before the hyperinflationary peak of February 1997. Due to the presence of mixed data series, we have to utilize the ARDL approach to cointegration. It is shown that inflation seems to be highly dependent on changes in the foreign exchange rate whereas monetary growth is not significant enough to explain inflation.

The complete lack of trust of economic agents in the Bulgarian banking system leads them to protect their savings by replacing Leva with foreign currencies. Consequently, the explosion in prices is not the result of money growth but rather a rejection of the domestic currency in favor of foreign currencies, above all the US Dollar and the Deutschmark. Besides, such a rejection is largely magnified due to the appearance of a price/exchange rate indexation mechanism as was the case during German hyperinflation.

From a theoretical point of view, such a result seems to argue in favor of the BP school in the case of a transition economy like Bulgaria in the 1990s. It also contains an important policy implication at least in some high inflation regimes. This empirical work suggests that high inflation should be brought under control by stabilizing the foreign exchange rate first instead of systematically focusing on monetary growth. In the longer run, countries experiencing high inflation and hyperinflation have no other viable solution than to put forward an industrial policy and conduct an import substitution strategy. Indeed, balance of payment issues are often involved because of an inadequate industrial sector and a very low ratio of investment to GDP (see, for example, Kulesza, 2017).

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